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WPS 0451

Country Economics Department
The World Bank
July 1990
WPS 451

Productivity, Imperfect Competition, and Trade Liberalization in Côte d'Ivoire

Ann E. Harrison

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Research on productivity often focuses on the relationship between productivity increases and such structural changes in an economy as trade reform.

If those structural changes affect the nature of competition or affect scale, however, both the changes and the level of change in productivity may be mismeasured.

Harrison extended previous studies to measure the relationship between productivity, market power, and trade reforms. Using a panel of 287 firms in Côte d'Ivoire, she analyzed changes in firm behavior and productivity, measuring market power before and after the 1985 trade reform.

Harrison found evidence that market power fell in several sectors following the changes in trade policy. She also shows that ignoring the effects of liberalization has led researchers to mismeasure the effect of trade reform on productivity.

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Productivity, Imperfect Competition, and Trade Liberalization in the Côte d'Ivoire

by
Ann E. Harrison*

Table of Contents

* This paper is a revised version of an earlier draft prepared in June 1989. I would like to thank David Card, Avinash Dixit, Gene Grossman, Zvi Griliches, Ramon Lopez, John Newman, Duncan Thomas, Jim Tybout, Jaime de Melo, and John Page for helpful comments and suggestions. This paper was prepared under funding from the World Bank research project "Industrial Competition, Productive Efficiency, and Their Relation to Trade Regimes (RPO 674-46).

Theoretical arguments for the gains from trade have traditionally rested on the concept of allocative efficiency. In an open economy with unrestricted trade, resources are more likely to be allocated in areas where a country has a comparative advantage. The recent emphasis on imperfectly competitive markets in international trade creates yet another argument for the welfare benefits of free trade: in a protected market dominated by several firms, trade reform will lead to increased competition. 1/

Despite the consensus on the theoretical benefits of free trade, the empirical evidence is often inconclusive. There is little evidence linking trade reform with increased competition, particularly for developing countries. Since most studies on trade reform use aggregate data across sectors or countries, these cannot capture changes in behavior at the individual firm level. 2/ Such studies typically analyze the effects of trade reform on behavior indirectly by addressing such issues as sectoral growth before and after the reforms.

Even more surprising is the lack of definitive evidence linking trade reform and productivity growth. Several recent overviews of the links between trade regimes and productivity growth (Bhagwati (1988), Havrylyshyn (1987), Nishimizu and Page (1987)) suggest that the evidence is mixed. As an illustration, we contrast Nishimizu and Robinson (1984) with Nishimizu and Page (1987). The earlier study finds a negative correlation between productivity growth and the degree of import substitution in four developing economies. The later study, which covers a different time period and a more extended set of countries, reverses the earlier finding.

One possible explanation for the lack of conclusive results may depend on how productivity is measured. The measurement of productivity pioneered by Solow (1957) has been used extensively to analyze technological change in both developing and developed countries. Solow derived a productivity measure, referred to as total or multi-factor productivity (TFP), which depends on the assumption that product markets are perfectly competitive. Yet shifts in trade policy are likely to alter the competitive environment, particularly in developing countries where domestic markets are often small and dominated by several firms.

Although the potential biases from assuming perfect competition have long been recognized (see Nishimizu (1979)), this paper implements a simple approach to correct TFP estimates for these biases. Recent papers by Robert Hall (1986,1988) and Domowitz, Hubbard, and Petersen (1988) derive a methodology which allows them to account for market power when estimating productivity. By relaxing the assumption that firms set price equal to marginal cost, Hall shows that previous estimates of productivity were spuriously procyclical. Domowitz, Hubbard and Petersen (1988) extend Hall's analysis to allow for material inputs and changing price-cost margins over time, then apply their methodology to aggregate industry data.

In this paper, we extend these earlier approaches to analyze changes in firm behavior and productivity during trade liberalization in the Cote d'Ivoire. For a panel of 287 firms, we estimate market power before and after a trade reform implemented in 1985. Our results suggest that price-cost margins fell in a number of sectors following the reform. However, since the reform was accompanied by a real appreciation in the exchange rate, part of the fall in margins was also due to the conjunction of the trade reform with the adverse exchange rate movement.

When productivity estimates are modified to account for changes in price-cost margins over the period, the positive correlation between trade reform and productivity is strengthened in some sectors and reversed in others. These results suggest that conclusions based on traditional productivity estimates are extremely sensitive to the assumptions about firm behavior. Section I outlines the theoretical approach and shows how ignoring the effects of liberalization on competition may lead researchers to mismeasure the effect of trade reform on productivity. Section II discusses trade policy changes in the Cote d'Ivoire and briefly describes the data. We present estimation results in Section III. Section IV explores the sensitivity of productivity measures to alternative specifications, including the possibility that the technology is characterized by increasing returns to scale. Finally, in Section V we incorporate our findings on market power to derive modified TFP estimates.

I. The Bias in Productivity Measurement

Our framework extends Hall (1986,1988) and Domowitz, Hubbard and Petersen (1988). We begin with a production function for firm i in industry j at time t:

$$Y_{ijt} = A_{jt}\, e^{f_{it}}\, G(L_{ijt}, K_{ijt}, M_{ijt}) \tag{1}$$

Output Y_{ijt} is produced by firm i with inputs labor L, capital K, and materials M. A_{jt} is an industry-specific index of Hicks-neutral technical progress, while f_{it} is a firm-specific parameter which allows for differences in firm technology. In our estimation, we will want to identify industry-wide productivity A. Totally differentiating (1), and dividing through by Y, we have

$$\frac{dY}{Y}_{ijt} = \frac{\partial Y}{\partial L}\frac{dL}{Y}_{ijt} + \frac{\partial Y}{\partial K}\frac{dK}{Y}_{ijt} + \frac{\partial Y}{\partial M}\frac{dM}{Y}_{ijt} + \frac{dA}{A}_{jt} + f_i \tag{2}$$

The element of imperfect competition enters (2) because firms with market power do not set the value marginal product $P(\partial Y/\partial L)$ equal to the factor price. If we assume Cournot behavior by firms, then we can derive the first order conditions from each firm's profit maximization and write each of the partial derivatives $\partial Y/\partial L$, $\partial Y/\partial K$, and $\partial Y/\partial M$ as follows:

$$\partial Y/\partial M_{ijt} = \frac{w}{P}_{jt} \left[\frac{1}{1 + (S_{ij}/e_j)}\right] = \frac{w}{P}_{jt}\,\mu_{ij} \tag{3a}$$

$$\partial Y/\partial M_{ijt} = \frac{r}{P}_{jt} \left[\frac{1}{1 + (S_{ij}/e_j)}\right] = \frac{r}{P}_{jt}\,\mu_{ij} \tag{3b}$$

$$\partial Y/\partial M_{ijt} = \frac{n}{P}_{jt} \left[\frac{1}{1 + (S_{ij}/e_j)}\right] = \frac{n}{P}_{jt}\,\mu_{ij} \tag{3c}$$

Factor prices are given by w (the wage), r (the rental cost of capital), and n (the price of material inputs). If firm i is not perfectly competitive, then the value of the marginal product exceeds the factor cost by some mark-up μ. Note that μ is a function of the industry demand elasticity e (which is negative) and the firm's market share S. We assume that firms in the same industry face the same e_j, and that the demand elasticity is constant over the time period estimated. Initially, we also assume that $S_{ijt} = S_j$ for all i and t. This implies that firm shares (by industry) are relatively stable over the period of estimation and of equal size. We relax this assumption in Section IV, allowing shares (and consequently mark-ups) to vary over firms and over time. The above specification also imposes the restriction that the mark-up must always be greater than or equal to unity. In a one-period oligopoly model, firms are not allowed to make short-run losses because such behavior would not be rational. This restriction is probably unnecessary but is not rejected by the data (see Section III).

Substituting 3a-3c into (2) and rearranging terms, we have

$$\frac{dY}{Y}_{ijt} = \mu_j \left[\frac{wL}{PY}\frac{dL}{L} + \frac{rK}{PY}\frac{dK}{K} + \frac{nM}{PY}\frac{dM}{M} \right]_{ijt} + \frac{dA}{A}_{jt} + f_i \qquad (4)$$

The value of wL/PY, rK/PY and nM/PY is simply the share of each factor (labor, capital, materials) in total output. We shall denote the share of labor and materials as α_l and α_m. Under constant returns to scale, the factor shares would sum to $1/\mu$, but we will retain a general formulation and allow the sum of the factor shares to equal β/μ, where β may be less than or greater than one.3/ Rewriting (4),

$$dy_{ijt} = \mu_j \; [\, \alpha_l dl + \alpha_m dm \,]_{ijt} + (\beta-1)_j \, dK/K_{ijt} + dA/A_{jt} + f_i \qquad (5)$$

Lower case variables y, l and m are equal to ln(Y/K), ln(L/K), and ln(M/K). The mark-up μ is just the coefficient on the changes in L/K and M/K, weighted by their respective shares in output. The specification in (5) imposes the restriction that the markup coefficient is equal across labor and material inputs. We test this restriction in Section III.

To see how estimates of productivity change dA/A could be biased due to the presence of imperfect competition, for the moment we will assume constant returns to scale (β=1), ignore the firm-specific effect, and rewrite (5) as

$$dy - \alpha_l dl - \alpha_m dm = \phi = (\mu-1)(\alpha_l dl + \alpha_m dm) + dA/A \qquad (6)$$

We will refer to ϕ as the "observed" productivity measure, and dA/A as the "true" productivity change. Under perfect competition , μ=1 and ϕ = dA/A. The Solow measure of productivity dA/A is unbiased.

If μ is greater than 1, however, there are two possible sources of bias. First, we may get bias in estimating the <u>level</u> of productivity change dA/A. If l and m are rising (falling), then dA/A is over (under) estimated. Second, <u>changes</u> in the trend rate of growth of productivity will be mismeasured. Figure 1 outlines the possible biases in estimating changes in the trend rate of growth of productivity. As an example, we explore the case where price-cost margins

Figure 1

Direction of Bias in Productivity Estimates

dA/A = true productivity
ϕ = observed productivity

	Case A	Case B
	$\alpha_l dl - \alpha_m dm > 0$	$\alpha_l dl - \alpha_m dm < 0$
Pre-reform		
$\mu > 1$	$\phi_1 > dA_1/A_1$	$\phi_1 < dA_1/A_1$
Post-reform		
$\mu = 1$	$\phi_2 = dA_2/A_2$	$\phi_2 = dA_2/A_2$
Net change	$\phi_2 - \phi_1 < dA_2/A_2 - dA_1/A_1$	$\phi_2 - \phi_1 > dA_2/A_2 - dA_1/A_1$
Bias	Productivity gains under-estimated	Productivity gains over-estimated

exceed one and firms have market power prior to a trade reform. In this case, the level of observed productivity ϕ will be greater (less) than the true measure if l and m are rising (falling). If the trade reform is accompanied by a fall in market power (possibly due to increases in the perceived elasticity of demand), price-cost margins fall to unity and measured productivity will equal the true productivity measure dA/A. However, if we are interested in comparing productivity before and after the changes in trade policy, we are likely to incorrectly assess the true change in dA/A. As illustrated in Figure 1, the direction of the bias cannot be predicted on the basis of (6): we are equally likely to overestimate or underestimate the increase in productivity following the reform.

To see how estimates of productivity dA/A could be biased due to increasing or decreasing returns to scale, let us assume perfect competition ($\mu = 1$) and rewrite (5) as (6)'

$$dy - \alpha_l dl - \alpha_m dm = \phi = (\beta-1)_j \; dK/K \; + dA/A \qquad (6)'$$

If β exceeds 1, then the technology is characterized by increasing returns. Observed productivity measure will exceed the true value dA/A as long as dK/K is positive, so TFP is over-estimated. Under decreasing returns, TFP will be under-estimated. Since increasing returns are consistent with imperfect competition, empirically we should observe b greater than 1 and market power concurrently.

II. Trade Policy in the Cote d'Ivoire

The trade regime in the Cote d'Ivoire became increasingly restrictive in the 1970s. In 1973, a major restructuring of the tariff code increased nominal tariff rates and raised levels of effective protection by implementing an escalated tariff structure. In the second half of the 1970s and in the early 1980s, quantitative restrictions and arbitrary reference prices were introduced on a wide range of imports competing with domestic manufactures. Table 1 indicates the extent of trade protection across industrial sectors before the reform, as measured by effective protection coefficients and the number of import licences (quotas) issued across sectors. Textiles and food-related manufacturing received the highest effective protection, followed by chemicals. Quotas were also high in food processing and beverages, followed by textiles and chemicals.

During the boom years in the second half of the 1970s, the Cote d'Ivoire benefitted from the surge in world coffee and cocoa prices. The increases in revenue, most of which were captured by the government, were used to promote investment and expand public spending and infrastructure. The severe macroeconomic imbalances that followed the fall in coffee prices forced the government to adopt an austerity program in 1982. The adjustment program was followed by a major trade reform introduced in mid-1984.

The trade reform was implemented in 1985 and extended in 1986 and early 1987. The reform removed quantitative restrictions and reference prices, rationalized the tariff structure, and introduced temporary tariff surcharges. The goal of the tariff reform was to equalize effective protection across different sectors by lowering tariffs on final goods and raising tariffs on inputs and intermediate goods. The surcharges declined over a five-year period to allow firms previously protected by non-tariff measures to adjust. Tariff changes and the removal of quotas was implemented in two phases. In the first phase (1985), reforms were imposed on key sectors including textiles and food processing. In the second phase (late 1986, early 1987), the reform was extended to the rest of the manufacturing sector (fertilizers, machinery).

Cote d'Ivoire's nominal exchange rate is fixed in relation to the French franc at a rate which is the same for a number of franc zone African countries. When the french franc appreciated against the US dollar between 1985 and 1988, the Ivorian franc became considerably overvalued in real terms. Consequently, the reform was conducted in conjunction with an environment which lowered the competitiveness of exports on world markets. lthough the government simulated a partial devaluation through an export subsidy scheme for manufactured exports, the first subsidy payments were delayed until mid-1986 and payments were concentrated in several large firms. The government's inability to compensate exporting firms for the real appreciation meant that the export sector was adversely affected. Consequently, we should see a fall in price-cost margins for exporting sectors in the post trade reform period.

To account for changes in behavior and productivity during the trade reforms in Cote d'Ivoire, we will want to modify (5) to allow for a change in mark-ups by firms during the post-1985 period. Changes in behavior would be captured by adding an interactive slope dummy to dx in (5). If trade reform induced a shift in the overall level of productivity, then we should also include an intercept dummy. We then have the estimating equation:

$$dy_{ijt} = B_{1j} dx_{ijt} + B_{2j} [D\, dx]_{ijt} + B_{3j} D + B_{4j}\, dk_{ijt} + dA/A_{jt} + f_i \qquad (7)$$

where

$$B_{1j} = \mu_j$$
$$B_{4j} = (\beta - 1)_j$$
$$dx = [\alpha_l dl + \alpha_m dm]$$
$$dk = dK/K$$

D is 1 for 1985-87 and 0 otherwise. If trade policy changes did in fact lead to more competitive firm behavior, the coefficient B_2 on [D dx] should be negative, reflecting the fall in mark-ups when firms are exposed to international competition. The coefficient B_4 is equal to the scale parameter ß minus one. If the coefficient is greater than zero, the technology is characterized by increasing returns; if it is equal to zero, constant returns; if less than zero, decreasing returns. The productivity term dA/A can be thought of as the sum of a constant industry level rate B_{0j} plus a residual u_{it}. This yields our final estimating equation:

$$dy_{ijt} = B_{1j}dx_{ijt} + B_{2j}\,[D\,dx\,]_{ijt} + B_{3j}\,D + B_{4j}\,dk_{ijt} + B_{0j} + f_i + u_{it} \quad (7)$$

If the individual effect f_i is fixed, we can estimate (7) using a standard fixed effect approach. If, however, fi is random, then estimating (7) as a fixed effect model will yield consistent but inefficient estimates. Under a random effect model, the most efficient estimate of (7) requires generalized least squares (GLS). 4/ Since the individual effect in (1) is modelled as a difference in production technology which is not likely to vary randomly across individuals, the fixed effect model is probably the more appropriate specification.

Data

The firm data is taken from information sent to the Banque de Donnees Financieres (BdDF), which is instructed to gather annual information on all industrial firms. The number of firms in individual years ranges from around 250 in the mid-1970s to nearly 500 in the mid-1980s. Although the coverage of the industrial sector is incomplete (informal enterprises are excluded and small formal firms are under-represented), the BdDF covers almost all large and medium-size formal manufacturing enterprises. We chose our sample of 287 firms by selecting out those enterprises with a complete time series. Although we include firms which were only present during part of the 1975-1987 period, we exclude firms which had missing values between their entry and exit dates. Table 2 shows that our sample includes the major firms in each of the manufacturing subsectors. The sample firms accounted for over half of all sales in 1987 for all sectors covered. For 10 of the 13 sectors, these firms accounted for over 70 percent of all output in 1987.

We estimate (7) using our panel of 287 Ivorian firms during the period 1975 to 1987. Since firms in different sectors are likely to exhibit different degrees of competition and face different sets of demand elasticities, we estimate the equation separately across 9 sectors (see Table 2). The approach requires data on real output, capital stock, labor and material inputs, and the shares of labor and materials in total output. Total sales and material inputs were deflated by 2 digit sectoral level price deflators to obtain a real output and materials series. We also calculated a material input price deflator based on input-output tables for each of the sectors, but the estimation results were unaffected and are not reported here. Real capital stock was constructed in two steps. First, for those firms that reported across the entire sample period, we used the perpetual inventory method. Real capital stock in period t is defined as:

$$K_{it} = (1-d)K_{i,t-1} + I_t \qquad (8)$$

As a benchmark, we used 1976 capital stock for each firm and then added real investment while accounting for depreciation. Real investment was computed by deflating nominal investment by sector-specific investment price deflators. To construct a base year real capital stock for the remaining sample of firms which entered after 1976, we first constructed a capital stock price deflator (KPD) using data on firms that were present in all years:

$$KPD_{jt} = \frac{\sum_{i=1}^{n} K_{ijt}}{\sum_{i=1}^{n} NK_{ijt}} \tag{9}$$

KPD_{jt} is the capital stock deflator for sector j in year t. It was constructed using the ratio of the real capital stock computed in (8) to the nominal capital stock (NK) reported firms that were present in all years. The real base year capital stock for a firm entering in year t is then given by the product

$$K_{ijt} = (KPD_{jt})\ (NK_{it}) \tag{10}$$

where t is the base year capital stock for firm i. For subsequent years, real capital stock is then computed using equation (8).

The total number of employees for each firm was used as a measure of labor input. The dataset does not include hours worked, which is the variable used by Hall (1988) and Domowitz, Hubbard and Petersen (1988) to measure labor input. However, using numbers of employees rather than hours should be accurate as long as there has not been a trend in economy-wide hours per employee. The results of household surveys for the Cote d'Ivoire (the LSMS World Bank project) for 1985 and 1986 indicate that average hours worked per employee did not change significantly over this two year period. 5/ Since these two years include both a year of unusual growth (1985) as well as a recession (1986), the fact that hours worked per employee was relatively stable suggests that the biases in using numbers of employees should not be too important. However, we cannot dismiss the possibility that there may have been a trend-line change in hours over the longer 1976-87 period.

Since there are only several firms in some of the industries listed in Table 2, we aggregated our firm sample into nine sectors: grain processing, food processing, other food, textiles, chemicals, transport, machinery, wood, and paper products. Sample means by sector for the pre- and post- trade reform periods are given in Table 3. Growth rates dropped in 6 of the 9 sectors, but the average annual growth rate stayed constant, averaging 4.8 percent. Over the 1976-1984 period, the average growth rate is high due to the boom in the economy in the 1970s. When trade reforms were introduced in 1985, the economy experienced a period of growth, but 1986 and particularly 1987 was a recessionary period. The burden of the adjustment appears to have fallen disproportionately on the labor force, with the annual average growth of employees falling from 2.0 percent to 1.4 percent. One shortcoming of the labor input variable is that it measures the number of permanent employees hired by the firm, but does not include information on temporary workers. One possibility, which we do not investigate, is that the fall in number of permanent employees was accompanied by an increase in the temporary labor force. However, the fall in employment by the formal sector, documented in Table 2, has been confirmed by other studies. 6/

In Table 3, we also report total factor productivity, unadjusted for market power effects. The productivity measure was calculated using a Tornquist index number formula (see definition in Table 3). Under trade reform, the unadjusted measure shows productivity increases in most sectors (food processing, chemicals, transport, machinery, and paper products) but declines in others (textiles and wood products). On average productivity growth accelerated under the trade reform, rising from 0.7 percent to 1.5 percent.

III. Estimation

We first estimate (7) using ordinary least squares (OLS) and adopting the assumption of constant returns to scale. We also estimate (7) assuming alternately a fixed effect and a random effect specification. Since inputs and output are jointly determined, however, we then estimate (7) using an instrumental variables technique. In Section IV we will explore the consequences of relaxing the assumptions of constant returns and equal mark-ups across firms.

The OLS results, without accounting for fixed effects, are presented in Table 4. The coefficient B_1 should measure the extent of market power across sectors, while B_2 indicates the change in price-cost margins under trade reform. The mark-up of price over marginal cost is highest in the food and textiles sectors, ranging from 28 percent for the "other food" category to 13 percent for textiles. We note that these are the sectors with the highest levels of protection prior to trade reform (refer to Table 1). At the same time, these are also the sectors with the greatest degree of outward orientation (See Table A.1). Since we cannot separate production for export and the domestic market due to the fact that inputs are not recorded separately, it is impossible to test the hypothesis that exporters charge high prices in the domestic market but price more competitively abroad.

The coefficient on B_2 is negative in six out of nine sectors, which would support the hypothesis that price-cost margins fell during the trade reforms. However, B_2 is only statistically significant (and negative) for the textile sector, which suggests that the changes in trade policy generally did not affect price-cost margins <u>except</u> in the textile sector, where the coefficient has the expected negative sign. Anecdotal evidence suggests that the conversion of quotas to tariffs led to large scale underinvoicing, which was particularly severe in the textile sector. For the other sectors, it is possible that changes made in 1987 in the structure of protection for chemicals and transport were too recent to show up in our data. The coefficient on B_3 indicates the change in productivity growth during the trade policy reforms. The coefficient is positive for 6 of the 9 sectors, but only statistically significant and positive for the paper sector. Paper products experienced an unusual increase in growth during the trade reform period (see Table 3). Since productivity is typically procyclical, the statistically significant increase in productivity growth during the reform period may only partly be attributed to changes in the trade regime.

We noted earlier that our specification imposes the restriction that mark-ups are equal across labor and material inputs. We test this restriction by allowing separate coefficients on labor and material inputs in (5). The F-value for the null hypothesis that the coefficients are equal is also included in Table 4. Equality of input coefficients is accepted for six out of nine sectors in our sample. Abbott, Griliches, and Hausman (1989) estimated a similar equation using US data for the cement industry and found that the restriction of equal coefficients was not accepted. Rotemberg and Summers (1988) suggest that failure of the model specification may reflect labor hoarding. We noted earlier that movements in labor inputs may not be fully reflected due to the possibility of hiring temporary workers. In our data, rejection of the specification test occurred when the coefficient on the labor input was close to zero, indicating no correlation between output and labor inputs. However, the coefficient should still capture price-cost margins because of movements in material inputs.

The specification above ignores any fixed effects which may be specific to the individual firms. If fixed effects are present, then our estimating equation is mis-specified and the coefficients may be biased. One way to test for this possibility is to compare the OLS results with estimates which allow for a firm-specific effect which is constant over time. We test for this alternative specification using a standard fixed-effect, within-group estimator. The within estimates are reported in Table 5. There is virtually no movement in the parameters. An F-test of the restriction that the OLS and fixed effect coefficients are equal is accepted for all sectors. Either a fixed effect is not present, or is removed by estimating the equation in growth rates instead of in levels. We also explore the possibility that the firm-specific effect f_i is random. If f_i is random and not fixed, then estimating a fixed effect model using OLS will be unbiased but not as efficient as generalized least squares estimation (GLS). The GLS estimates are presented in Appendix Table A.2. The random effects specification does not yield statistically different estimates from OLS.

The OLS estimates are likely to be biased since inputs and output are simultaneously determined by the firm. Table 6 presents instrumental variables (IV) estimates under the maintained assumption of constant returns to scale. The instruments should be correlated with the endogenous right-hand side variables dX and D dX but independent of any demand or productivity shocks. As instruments we use the second lag of the nominal exchange rate, a price index for energy, the second lags of employment and materials, and the second lag of the firm's debt to sales ratio. The price of energy should be correlated with input decisions but independent of any demand shocks or productivity shocks affecting the firm. We use the second lag (in levels) for inputs instead of the first lag since the right-hand side variables already include current and lagged values. We also include the second lag of the firm's debt to sales ratio, under the assumption that the firm's borrowings should be correlated with ability to expand inputs but are predetermined.

Following Bowden and Turkington (1984), we instrument the product D dX using a nonlinear combination of the dummy D and dX. In our case, this is just the set of variables D, instruments for dX, and the product of D and the instruments.7/ The instrumental variable estimates in Table 6 were tested for stability using various alternative sets of instruments, including a composite wage index. Our experience (which is confirmed by Abbot et al (1989)) suggests that the standard instruments which are used in these types of regressions, such as GNP, are likely to be correlated with the error term and may lead to biased estimates of price-cost margins. Alternative specifications which employed GNP as an instrument in the first stage of the regression led to rejection of the over-identification tests for the exogeneity of our instruments. 8/

The estimated coefficients in Table 6 show a similar pattern to the OLS estimates. Mark-ups are highest for food-related and textile firms. However, due to the larger standard errors, we can only reject the null hypothesis of perfect competition for one of the food sectors, machinery, and wood products. Mark-ups generally fall during the trade reform period, as indicated by the negative coefficient on B_2. The fall in mark-ups is statistically significant (and negative) for three sectors: food processing, textiles, and wood products. Since the level of protection in textiles was quite high before the trade reform, it is likely that the fall in quotas contributed to lower mark-ups. The fall in margins for food processing and wood products, however, seems to be linked to the adverse impact of the franc's appreciation. Table A.1 shows the share of exports in total sales for the firms in our sample. The food processing (cocoa, coffee) and wood sectors are the most export oriented of all nine sectors. Despite the subsidies to offset the negative impact of appreciation on exporters, it appears that the outward oriented sectors were most negatively affected during the trade reform.

Table 6 reports the F-value when we test the restriction that the price-cost margin cannot fall below one, which is an implication of our model. For all sectors, margins were either positive or not statitistically different from unity. We also test the validity of our instruments in Table 6. Newey (1985) suggests a chi-square test which may be applied if the estimating equation is overidentified and a subset of the instruments is assumed to be valid. A regression of the residuals from the first stage regression on the instruments yields a chi-square test of the validity of our instruments. The results of this test, shown in Table 6, suggest that in all cases our instruments are valid.

IV. Alternative Specifications

Increasing or Decreasing Returns to Scale

One important source of mis-specification arises from assuming that the technology is characterized by constant returns, which permits us to omit dk from equation (7). Recall from (7) that the coefficient on dk is given by β-1, where β is the returns to scale parameter. Under constant returns, β equals 1 and the coefficient on dk is zero. Under increasing returns, however, ß exceeds one and the coefficient on dk is positive. If the technology is not characterized by constant returns, then omitting dk induces a classic omitted variable bias. The extent of the bias is given by the product R (ß-1), where ß-1 is the coefficient on dk and R is the coefficient of dk regressed on dx (see Schmidt (1976)). Assuming that R is negative, price-cost margins will be under-estimated with increasing returns to scale. If, on the other hand, the technology exhibits decreasing returns, the margins are over-estimated.

We investigate the possibility of bias due to the technology assumptions in Table 7. Table 7 shows the same pattern of price-cost margins as earlier estimates. Margins are highest in textiles, export-related products (wood, food processing), and chemicals. The coefficient on dk is statistically significant for one product group, "other food". For this product, the mark-up falls when the capital stock is included in the regression equation. The negative coefficient on the capital stock variable, which indicates decreasing returns, means that price-cost margin are over-estimated for that sector. For all other sectors, however, the coefficient on the capital stock variable is not significant. For most of our sample, specification of the technology as characterized by constant returns is not inappropriate.

Another potential shortcoming of the original specification is the assumption, implicit in the first order conditions 3a-3c, that the value of the marginal product of capital is set equal to a mark up μ multiplied by the cost of capital (r) in every period. A profit-maximizing firm might choose labor (L) and materials (M) in the short run, and take capital (K) to be predetermined. It can be shown that if we alter the original maximization problem of the firm so that the production technology is given by

$$Y = A \quad G(L , M ; K) \tag{1'}$$

then equation (5) becomes

$$dy = \mu \left[\alpha_1 dl + \alpha_m dm \right] + \left[\mu \ (\alpha_1 dl + \alpha_m dm) - 1 + \delta \right] dK/K + dA/A \tag{5'}$$

$$\text{where } \delta = \partial G/\partial K \ (K/Y)$$

Empirically, we cannot distinguish between equation (5)' above and the specification which allows for increasing (or decreasing) returns to scale. Both theoretical models suggest that dK/K should be included on the right hand side of the estimating equation. However, since our results indicate that in general dK/K is insignificant, the fact that we cannot assign a specific interpretation to the coefficient on dK/K is not particularly important in this paper. However, any approach which seeks to interpret the coefficient on dK/K as an indication of scale economies should take into consideration the dual interpretations of the parameter. A significant coefficient on dK/K may indicate either non-contant returns to scale or short-run profit maximization which takes capital stock as given.

Varying Price-cost margins across firms and over time

Another possible source of mis-specification arises from the possibility that the price-cost margin should vary across firms and over time. The justification for firm-specific and time varying shares comes from equations 3(a)-3(c), which show that the mark-up is related to firm shares through the equation:

$$\mu(S)_{ijt} = \frac{1}{1 + [S_{ijt}/e_j]} = \frac{e}{e + S_{ijt}} \qquad (11)$$

Recall that the elasticity e is negative. Mark-ups in the simple Cournot framework should be positively related to firm shares. To account for the dependence of μ on S, we take a first order Taylor approximation of (11) around an initial point a:

$$\mu(S) = \frac{e}{e + a} + \frac{-e\ (S - a)}{(e + a)^2}$$

Rearranging terms, we have

$$\mu(S) = B_1 + B_4\ S \qquad (12)$$

$$\text{where } B_1 = \frac{e}{(e + a)} + \frac{ea}{(e + a)^2}$$

$$B_5 = \frac{-\ e}{(e + a)^2}$$

Combining (7) and (12),

$$dy_{ijt} = B_{0j} + B_{1j}dx_{ijt} + B_{2j}\ [D\ dx]_{ijt} + B_{3j}\ D + \\ + B4_j\ [S\ dx]_{ijt} + B_{5j}\ [S\ D\ dx]_{ijt} + u_{it} \qquad (13)$$

If the price-cost margin does not vary across firms, then the coefficients B4 and B5 should be statistically insignificant and the coefficient on dx collapses to the mark-up parameter u. We calculate firm shares by using as a denominator total output by sector less export sales, available from the BdDF database. The firm's share of domestic output is the ratio of the firm's domestic sales less export sales to total sector (domestic) output. One of the shortcomings of this approach is the lack of import data. Consequently, the variable S does not represent the firm's share of total domestic consumption, but it does indicate the firm's share of domestically produced output.

In Table 8, our estimates for (13) show that the coefficient on S is statistically insignificant across all sectors. The statistical insignificance of S seems to indicate that differences in shares, either across firms and over time, are not a source of varying mark-ups. The patterns observed earlier are also exhibited in Table 8. Price-cost margins are highest for textiles and export oriented sectors. Nevertheless, there are a number of problems with this approach. The coefficient on S should be positive (see equation 12), yet for half the sectors it is in fact negative. One possibility is that a simple Cournot model may not be appropriate as an explanation for the observed mark-ups. Another problem is that firm shares are calculated as a fraction of total domestic output. Finally, the estimates may be highly imprecise since nearly all the right-hand side variables are endogenous. Nevertheless, our instrumental variables estimates in Table 8 do give statistically significant coefficients for the price-cost mark-up, in contrast to the insignificant estimates on S. In future work, it may be desirable to explore the dependence of the price-cost margin on other (possibly firm-specific) factors.

Changes in Capacity Utilization

Another possible source of misspecification arises from the fact that observed price-cost margins may fluctuate as capacity utilization changes over the business cycle. If, for example, trade reform was accompanied by contractions in output of the tradeable sectors, then changes in mark-ups may reflect capacity changes rather than shifts in competitive behavior. In this case, we only observe measured capital stock K^*. The true K is equal to K^*E, where E reflects changes in utilization of capacity.

If the production function is given by $Y = A\ g(L, K^*E, M)$ then the resulting estimating equation becomes

$$dy - de = B_0 + B_1 \{ \alpha_l dl + \alpha_m dm - (\alpha_l + \alpha_m)\ de \}$$
$$+ B2\ [D \{ \alpha_l dl + \alpha_m dm - (\alpha_l + \alpha_m)\ de \}] + B_3 D \qquad (14)$$

Since we do not have estimates of capacity utilization at the firm level, we employ a measure of total energy use as a proxy. A plant's energy use is the input component most likely to vary as capacity utilization fluctuates. The OLS and instrumental variable estimates for equation (14) are shown on Table 9. The general patterns observed earlier are reproduced again in Table 9. In addition, the fall in margins becomes even more significant following the trade reform, for both the OLS and the instrumental variable estimates.

V. Modified TFP Estimates

In Section I we showed that TFP can be mismeasured in the presence of either imperfect competition or non-constant returns to scale. The results from Section III indicate that market power cannot be rejected for a number of manufacturing sectors in the Cote d'Ivoire. Here we incorporate those findings to analyze productivity before and after the trade reforms.

Although productivity may be estimated in a number of different ways, one standard approach is to use the Tornquist index number formula, which is a discrete approximation to the formula derived in equation (6):

$$TFP = [\ln Y_t - \ln Y_{t-1}] - [\alpha_l (\ln L_t - \ln L_{t-1}) + \alpha_m (\ln M_t - \ln M_{t-1}) + (1 - \alpha_l - \alpha_m)(\ln K_t - \ln K_{t-1})] \tag{15}$$

$$\text{where } \alpha_l = 1/2\ (\alpha_{lt} + \alpha_{lt-1})$$

$$\alpha_m = 1/2\ (\alpha_{mt} + \alpha_{mt-1})$$

If we incorporate the mark-up factor μ, equation (15) can be written as:

$$TFP = [\ln Y_t - \ln Y_{t-1}] - \mu\ [\alpha_l (\ln L_t - \ln L_{t-1}) + \alpha_m (\ln M_t - \ln M_{t-1}) + (1/\mu - \alpha_l - \alpha_m)(\ln K_t - \ln K_{t-1})] \tag{16}$$

Estimates of μ were taken from Table 9 to calculate revised TFP estimates before and under the trade reforms. Estimates of productivity using both the original and revised definitions (equations (15) and (16) respectively) are presented in Table 10. Since the model imposes the restriction that the price-cost margin must be greater than or equal to unity, we impose the restriction that margins cannot fall below unity. The original estimates are reproduced from Table 3. Under the assumption of perfect competition, we find that productivity increased under trade reform for six of the nine sectors. Over all sectors, productivity rose from an average of .7 to 1.5 percent annually. If we relax the assumption of perfect competition, the gain in productivity is much smaller. Productivity only rises in 5 sectors, with smaller increases overall. Since labor and material inputs per unit of capital experienced negative growth prior to 1985, TFP is underestimated in equation (15), and the change in productivity after 1985 is overstated. On average, the adjusted productivity measure only rises slightly, from .7 to .9 percent growth annually. The results in Table 10 suggest that productivity estimates are highly sensitive to the assumption of perfect competition. For example, in the food sector when we introduce imperfect competition, the gain in productivity post-reform rises from .8 percent to 2 percent annually. On average, our data suggests that when we incorporate imperfect competition into the productivity estimates, there is no apparent relationship between productivity and trade reform. However, two aspects of the reform must be acknowledged. First, the estimated time period for the sample post reform is only three years. If there are any productivity gains associated with trade reform, such gains may only appear over a longer time period. Second, the reform was accompanied by a real appreciation of the currency, which adversely affected exporting sectors. Table 10 shows that productivity fell primarily in exporting sectors, but generally increased in other sectors.

Conclusion

Research on productivity has often focused on the relationship between productivity increases and structural changes in an economy, such as trade policy reform. If, however, those structural changes affect the nature of competition or have scale effects, then both the levels and the changes in productivity may be mismeasured. In this paper, we extend previous studies to measure the relationship between productivity, market power, and trade reform.

Using a panel of 287 firms in the Cote d'Ivoire, we test for imperfect competition before and after the 1985 trade reform. We find that protected sectors such as textiles have significant mark-ups of price over marginal cost. We also find evidence that price-cost margins fell between 1985 and 1987. However, the fall in mark-ups for exporting sectors is likely to be linked more to the real appreciation of the currency than to increases in domestic competition.

If we incorporate measures of price-cost margins into estimates of productivity, we find that these estimates are extremely sensitive to the usual assumptions made about competition. Whereas there seems to be a strong relationship between trade reform and productivity when we assume perfect competition in the product markets, this relationship almost disappears when we allow for varying mark-ups. These results may be qualified, however, by noting that when the exporting sectors are excluded from the analysis, there are productivity gains associated with trade reform. The results support recent arguments (see Rodrik (1988)) that the theoretical basis for a positive link between trade reform and productivity growth should be explored in future research. More analytical efforts are needed which explicitly model the possible links between trade policy and productivity growth.

Notes

1/ For an overview of this literature, see Helpman and Krugman (1989).

2/ An exception is de Melo and Urata (1986), which compares reported price-cost margins for two census years before and after reforms in Chile. Research on developed country data includes Domowitz, Hubbard, and Petersen (1986), who use aggregate data to find a negative relationship between import penetration and reported price-cost margins.

3/ To see why this holds, say we have a production function given by $Y = AL^{a}M^{b}K^{c}$, where a + b + c sum to ß, the scale parameter. If we take logs and differentiate, we see that

$$\frac{dY}{dL}\frac{L}{Y} + \frac{dY}{dM}\frac{M}{Y} + \frac{dY}{dK}\frac{K}{Y} = a + b + c = \beta$$

But from our first order conditions, $\frac{dY}{dL}\frac{L}{Y} = \mu\, \alpha_l$, etc. so

we have $\mu\, \alpha_l + \mu\, \alpha_m + \mu\, \alpha_k = \beta$.

4/ For a description of GLS for panel data, see Hsiao (1986).

5/ A regression of hours worked on several control variables (sex, age, location, education) and a year dummy yielded an insignificant coefficient on the year dummy with a t-statistic of 1.54.

6/ See "Cote d'Ivoire: Industrial Competitiveness During Economic Crisis and Adjustment", Klaus Lorch, 1989.

7/ If instead we had regressed dx on a set of instruments to obtain a predicted value for dx, and then calculated the predicted dx multiplied by D1, this would have yielded a biased coefficient on the product dx D1.

8/ Estimates of the alternative specifications available from the author on request.

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TABLE 1

Estimates of Protection Across Sectors, 1980 and 1982

	Sector	Nominal Tariff Protection Coefficient 1980	Effective Tariff Protection Coefficient 1980	Occurences of Import Licenses 1982
I	Food Processing and Beverages			
	Grain processing	-	-	6
	Food processing	1.359	3.340	34
	Beverages	-	-	17
	Oils	1.405	1.829	2
	Other	1.493	1.829	0
II	Textiles			
	Clothing	1.422	2.455	182
	Leather, footwear	1.303	1.307	39
III	Chemicals			
	Chemicals	1.309	1.337	52
	Rubber	1.387	1.489	10
IV	Transport, Machinery			
	Transport	1.245	1.245	17
	Machinery	1.230	1.230	9
V	Wood, paper products			
	Wood	1.258	1.276	4
	Paper	-	-	-

Source: "The Côte d'Ivoire in Transition: From Structural Adjustment to Self-sustained Growth," World Bank, 1987

TABLE 2

Sample Coverage

	Sector	Number of Firms	Percentage of Formal Manufacturing Sector 1987
I	Food Processing and Beverages	50	
	Grain processing	17	72%
	Food processing	15	91
	Beverages	5	98
	Oils	5	98
	Other	8	50
II	Textiles	39	
	Clothing	23	86
	Leather, footwear	16	53
III	Chemicals	48	
	Chemicals	43	76
	Rubber	5	98
IV	Transport, Machinery	81	
	Transport	28	68
	Machinery	53	74
V	Wood, paper products	69	
	Wood	45	76
	Paper	24	75
	Total firms	287	

TABLE 3

Sample Means of Growth Rates for Selected Variables

Sector	1976-1984					1985-1987				
	Real Output	Employment	Materials	Capital	TFP 1/	Real Output	Employment	Materials	Capital	TFP 1/
Grain processing	3.3	3.0	3.0	2.7	0.1	3.2	-9.9	5.5	0.2	1.4
Food processing	5.1	6.5	2.9	3.0	1.4	-2.7	-0.2	-3.7	5.4	0.9
Other food	10.9	7.7	10.2	8.7	0.5	4.0	4.3	2.9	2.7	1.3
Textiles	2.7	1.5	1.7	3.2	0.4	-0.5	1.1	-0.5	7.7	-1.5
Chemicals 2/	7.4	2.7	6.6	9.3	1.3	12.7	6.3	12.6	3.1	3.6
Transport	-2.3	-1.4	-2.2	4.8	-1.3	11.4	2.4	11.8	6.9	1.7
Machinery 2/	6.8	-0.7	6.9	7.9	1.6	-2.8	-4.6	-10.1	-2.8	5.3
Wood products	1.5	1.5	1.1	1.3	0.6	-0.4	2.3	3.3	2.1	-3.4
Paper products	4.3	2.9	4.8	4.4	-0.3	9.6	-0.1	4.4	4.5	6.2
All Sectors	4.8	2.0	4.3	5.6	0.7	4.9	1.4	4.5	3.7	1.5

1/ TFP is defined using the Tornquist index number formula, with
$TFP = [\ln Y_t - \ln Y_{t-1}] - [\bar{\alpha}_l (\ln L_t - \ln L_{t-1}) + \bar{\alpha}_m (\ln M_t - \ln M_{t-1}) + (1 - \bar{\alpha}_l - \bar{\alpha}_m) (\ln K_t - \ln K_{t-1})]$.
Labor and material shares are defined as follows: $\bar{\alpha}_l = 1/2\ (\alpha_{l_t} + \alpha_{l_{t-1}})$, $\bar{\alpha}_m = 1/2\ (\alpha_{m_t} + \alpha_{m_{t-1}})$.

2/ The two periods for parts of this sector are divided into 1976-86 and 1987.

TABLE 4

OLS Results

Estimating Equation: $dy = B_0 + B_1 (\alpha_1 dl + \alpha_m dm) + B_2[D \cdot (\alpha_1 dl + \alpha_m dm)] + B_3 \cdot D$

Sector	B_1	B_2	B_3	F-Value for B_1=1	F-Value for Input Coefficient Equal 1/	N	R^2
Grain processing	1.206 (.052)	-.042 (.082)	.004 (.023)	15.7	1.4	117	.88
Food processing	.911 (.061)	-.114 (.123)	-.015 (.037)	2.1	20.3	110	.73
Other food	1.281 (.051)	.065 (.142)	.013 (.027)	30.1	0.1	162	.82
Textiles	1.136 (.055)	-.125 (.086)	-.035 (.028)	6.1	0.0	260	.72
Chemicals	1.068 (.036)	-.029 (.070)	.022 (.019)	3.5	2.3	361	.77
Transport	1.027 (.038)	.013 (.079)	.021 (.023)	0.5	4.2	160	.86
Machinery	1.078 (.027)	-.001 (.202)	.040 (.035)	8.2	0.3	333	.83
Wood products	1.055 (.029)	-.051 (.070)	-.033 (.025)	3.6	44.1	284	.85
Paper products	1.055 (.056)	.179 (.087)	.077 (.026)	1.0	2.2	157	.82
All Sectors	1.078 (.014)	-.019 (.030)	.008 (.009)	33.0	32.3	1944	.80

1/ We estimate the modified equation $dy = B_1(\alpha_1 dl) + B_2(\alpha_m dm) + B_3[D \cdot \alpha_1 dl] + B_4[D \cdot \alpha_m dm) + B_5 \cdot D$. We then test the joint restriction that $B_1 = B_2$ and $B_3 = B_4$. The F-Value of the test is reported here.

TABLE 5

Comparison of OLS and within Estimates[1/]

Estimating Equation: $dy = B_0 + B_1\ (\alpha_1 dl + \alpha_m dm) + B_2[D \cdot (\alpha_1 dl + \alpha_m dm)] + B_3 \cdot D$

	B_1		B_2		
Sector	OLS	Within Estimate	OLS	Within Estimate	F-Value[2/]
Grain processing	1.206 (.052)	1.212 (.052)	-.042 (.082)	-.015 (.090)	0.0
Food processing	.911 (.061)	0.86 (.069)	-.114 (.123)	-.083 (.141)	0.1
Other food	1.281 (.051)	1.277 (.053)	.065 (.142)	.027 (.150)	0.0
Textiles	1.136 (.055)	1.137 (.058)	-.125 (.086)	-.139 (.090)	0.0
Chemicals	1.068 (.036)	1.044 (.037)	-.029 (.070)	-.021 (.073)	0.0
Transport	1.027 (.038)	1.038 (.041)	.013 (.079)	.022 (.083)	0.5
Machinery	1.078 (.027)	1.008 (.026)	-.001 (.202)	-.027 (.198)	0.7
Wood products	1.055 (.029)	1.039 (.030)	-.051 (.070)	-.100 (.077)	0.0
Paper products	1.055 (.056)	1.072 (.056)	.179 (.087)	.146 (.087)	0.1
All Sectors	1.078 (.014)	1.053 (.014)	-.019 (.030)	-.014 (.031)	0.4

1/ Within estimates are computed by taking the deviations from firm means over time for all variables.

2/ Tests the restriction that the coefficients B_1 and B_2 are equal for the OLS and within estimates.

TABLE 6

Instrumental Variables Estimates 1/

Estimating Equation: $dy = B_0 + B_1 (\alpha_l dl + \alpha_m dm) + B_2[D \cdot (\alpha_l dl + \alpha_m dm)] + B_3 \cdot D$

Sector	B_1 1/	B_2 1/	F-Value for B_1=1	F-Value for $B_1 + B_2$=1	Overid Test 2/
Grain processing	0.831 (.305)	.444 (.338)	0.3	3.6	8.2
Food processing	1.161 (.189)	-.947 (.480)	0.7	0.7	1.2
Other food	1.349 (.139)	-.249 (.319)	6.3	0.1	14.1
Textiles	1.333 (.252)	-.487 (.285)	1.8	1.3	3.8
Chemicals	1.299 (.322)	.157 (.417)	0.9	3.0	3.6
Transport	.816 (.150)	.158 (.202)	1.5	0.0	7.4
Machinery 3/	1.251 (.154)	-.199 (.640)	2.6	0.0	1.5
Wood products	1.485 (.205)	-.680 (.348)	5.6	0.4	1.5
Paper products	1.004 (.306)	.339 (.334)	0.0	6.4	8.6
All Sectors	1.241 (.203)	-.305 (.232)	1.4	0.3	1.1

1/ Instruments are D; the second lag of the nominal exchange rate; price index for energy; the second lags of employment, materials, and the share of debt in sales; and D interacted with these 5 variables.

2/ The overidentification test gives the chi-square statistic for the hypothesis that the instruments are accepted as valid. The critical 5% value of the chi-square (9) = 16.9. A higher value indicates rejection of the test.

3/ Instrument list excludes interaction of D and 1) the price index for energy, 2) the second lag of the nominal exchange rate.

TABLE 7

IV Estimates with no Restrictions on the Scale Technology

Estimating Equation: $dy = B_0 + B_1\, dx + B_2[D \cdot dx] + B_3 \cdot D + B_4 \cdot dk$ 1/

Sector	B_1 2/	B_2 2/	B_4 2/	Overid Test 3/
Grain processing	.968	.410	.147	7.9
	(.408)	(.346)	(.290)	
Food Processing	1.194	-.930	.085	1.1
	(.254)	(.494)	(.423)	
Other food	.856	-.316	-.522	12.1
	(.367)	(.321)	(.360)	
Textiles	1.468	-.513	.243	2.3
	(.206)	(.453)	(.427)	
Chemicals	1.306	.124	-.093	3.3
	(.327)	(.458)	(.511)	
Transport	.720	.203	-.159	6.4
	(.193)	(.215)	(.191)	
Machinery	1.138	.025	-.200	1.0
	(.214)	(.699)	(.268)	
Wood products	1.468	-.513	.243	3.3
	(.206)	(.453)	(.427)	
Paper products	.672	.450	-.332	8.2
	(.482)	(.352)	(.377)	
All sectors	1.314	-.353	.065	10.7
	(.417)	(.336)	(.315)	

1/ $dx = \alpha_1 dl + \alpha_m dm$
$dk = dK/K$

2/ See Table 6, footnotes 1/ and 3/, for instrument list.

3/ The overidentification test gives the chi-square statistic for the hypothesis that the instruments are accepted as valid. The critical 5% value of the chi-square (8) = 15.5.

TABLE 8

IV Estimates Allowing Mark-ups to Vary Across Firms and Over Time[1]

Estimating Equations:

(1) $dy = B_0 + B_1 dx + B_2[D \cdot dx] + B_3 \cdot D + B_4[S \cdot dx]$

(2) $dy = B_0 + B_1 dx + B_2[D \cdot dx] + B_3 \cdot D + B_4[S \cdot dx] + B_5[S \cdot D \cdot dx]$

	(1)		(2)		
	B_1	B_4	B_1	B_4	B_5
Grain processing	.866	-.697	.833	-.0[illegible]	-.940
	(.329)	(1.090)	(.335)	(1.935)	(2.327)
Food processing	1.212	-1.475	1.303	-4.092	3.986
	(.234)	(3.644)	(.299)	(6.355)	(7.843)
Other food	1.352	-.022	1.418	-.491	1.425
	(.229)	(1.293)	(.267)	(1.606)	(2.800)
Textiles	1.378	-.727	1.405	-1.152	1.329
	(.309)	(2.847)	(.334)	(3.477)	(6.154)
Chemicals	1.299	-.275	1.299	-.930	.738
	(.323)	(2.098)	(.322)	(6.240)	(6.623)
Transport	.822	-.377	.767	3.025	-4.084
	(.171)	(4.929)	(.250)	(11.990)	(13.215)
Machinery	1.244	.270	1.052	7.849	-42.189
	(.617)	(24.152)	(.756)	(29.618)	(70.081)
Wood products	1.641	-6.045	1.631	-5.639	-1.382
	(.255)	(5.409)	(.270)	(6.426)	(11.864)
Paper products	1.229	-2.332	1.199	-2.023	-.323
	(.388)	(1.946)	(.975)	(9.478)	(9.684)
All Sectors	1.116	1.441	1.186	.634	2.780
	(.279)	(2.224)	(.318)	(2.753)	(5.111)

1/ See table 6 for instrument list.

TABLE 9

OLS and IV Estimates Adjusting for Capacity Utilization

Estimating Equations:

(1) $dy = B_0 + B_1\,(\alpha_1 dl + \alpha_m d_m) + B_2[D\cdot(\alpha_1 dl + \alpha_m d_m)] + B_3\cdot D$

(2) $dy - de = B_0 + B_1\,(\alpha_1 dl + \alpha_m d_m - (\alpha_1+\alpha_m)\,de)$
$+ B_2[D\cdot(\alpha_1 dl + \alpha_m d_m - (\alpha_1+\alpha_m)\,de)] + B_3\cdot D$

	OLS			IV 1/	
Sector	(1)	(2)		(2)	
	B_1	B_1	B_2	B_1	B_2
Grain processing	1.222	1.173	.009	1.081	.220
	(.081)	(.083)	(.104)	(.195)	(.235)
Food processing	.859	.970	-.222	1.243	-.694
	(.091)	(.070)	(.120)	(.329)	(.423)
Other food	1.268	1.346	.054	1.741	-.474
	(.048)	(.043)	(.087)	(.193)	(.304)
Textiles	1.119	1.082	-.186	1.165	-.439
	(.067)	(.046)	(.081)	(.188)	(.268)
Chemicals	.996	1.050	.039	.878	.597
	(.053)	(.035)	(.064)	(.195)	(.355)
Transport	1.059	1.108	.009	1.133	-.081
	(.061)	(.049)	(.067)	(.143)	(.189)
Machinery	.964	.952	.043	1.168	-.748
	(.032)	(.027)	(.185)	(.226)	(.795)
Wood products	1.167	1.152	.022	1.226	-.608
	(.047)	(.046)	(.065)	(.169)	(.305)
Paper products	.939	.996	.254	.653	.807
	(.086)	(.061)	(.090)	(.431)	(.458)
All sectors	1.058	1.060	.048	1.362	-.399
	(.019)	(.015)	(.027)	(.181)	(.215)

1/ See table 6, footnotes 1/ and 3/ for instrument list.

TABLE 10

Sensitivity of TFP Estimates to Assumption of Perfect Competition [$\mu = 1$]

$$TFP = [\ln Y_t - \ln Y_{t-1}] - \mu\, [\bar{\alpha}_L (\ln L_t - \ln L_{t-1}) + \bar{\alpha}_M (\ln M_t - \ln M_{t-1}) + (1/\mu - \bar{\alpha}_M - \bar{\alpha}_L)(\ln K_t - \ln K_{t-1})]$$ 1/

Sector	TFP [μ=1]			TFP 2/ [$\mu \neq 1$]		
	Pre Reform	Post Reform	Dif-ference	Pre Reform	Post Reform	Dif-ference
Grain processing	0.1	1.4	1.3	0.04	0.92	0.88
Food processing	1.4	0.9	-0.5	1.2	0.9	-0.3
Other food	0.5	1.3	0.8	-0.7	1.3	2.0
Textiles	0.4	-1.5	-1.9	0.5	-1.5	-2.0
Chemicals	1.3	3.6	2.3	0.90	0.88	-0.02
Transport	-1.3	1.7	3.0	-.5	1.5	2.0
Machinery	1.6	5.3	4.0	2.1	5.3	3.2
Wood products	.6	-3.4	-4.0	0.7	-3.4	-4.1
Paper products	-0.3	6.2	6.5	-0.3	6.7	7.0
All sectors	0.7	1.5	0.8	0.7	0.9	0.2

1/ This is the Tornquist index number formula for TFP, modified to allow for imperfect competition and scale economies. Labor and material shares $\bar{\alpha}_L$ and $\bar{\alpha}_M$ are defined as follows: $\bar{\alpha}_L = 1/2\ (\alpha_{L_t} + \alpha_{L_{t-1}})$

$\bar{\alpha}_M = 1/2\ (\alpha_{M_t} + \alpha_{M_{t-1}})$

2\ Estimates for μ taken from Table 9.

TABLE A.1

Share of Exports in Total Sales

Sector	1976-1984 (%)	1985-1987 (%)
Grain processing	4.3	4.3
Food processing	69.5	72.6
Other food	8.0	9.0
Textiles	17.0	10.0
Chemicals	20.7	22.4
Transport	6.6	6.2
Machinery	3.9	3.7
Wood products	43.2	41.0
Paper products	2.4	5.2
All sectors	18.2	19.7

Table A.2

GLS Estimates 1/

Estimating Equation: $dy = B_0 + B_1 (\alpha_1 dl + \alpha_m dm) + B_2[D \cdot (\alpha_1 dl + \alpha_m dm)] + B_3 \cdot D$

Sector	B_1	B_2	F-Value 2/
Grain processing	1.208 (.052)	-.041 (.083)	0.0
Food processing	.905 (.062)	-.112 (.125)	0.0
Textiles	1.135 (.055)	-.122 (.086)	0.0
Chemicals	1.066 (.037)	-.029 (.071)	0.0
Transport	1.029 (.039)	.013 (.079)	0.0
Machinery	1.060 (.027)	-.019 (.203)	0.1
Wood products	1.053 (.029)	-.055 (.071)	0.0
Paper products	1.059 (.056)	.175 (.088)	0.0
All sectors	1.072 (.014)	-.016 (.030)	0.0

1/ GLS estimates are computed by taking the deviations from firm means over time for all variables, given by $dy_{it} - a\bar{y}_i$. The value of a is equal to $1 - \sqrt{\theta^2}$, where $\theta^2 = \dfrac{\sigma_\epsilon^2}{\sigma_\epsilon^2 + T_i \sigma_\alpha^2}$. σ_ϵ^2 and σ_α^2 are estimated from between and within residuals, and T_i is the number of observations for each ith firm.

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